The Decline of Welfare Benefits in the U.S.: The Role of Wage Inequality

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Abstract:
Welfare benefits in the U.S. have experienced a much-studied secular decline since the mid-1970s. We explore a new hypothesis for this decline related to the increase in wage inequality in the labor market and the decline of real wages at the bottom of the distribution: we posit that voters prefer benefits which are tied to low-skilled wages. We test the hypothesis using a 1969–1992 panel of state-level data. An additional contribution of our analysis is the use of General Social Survey data on voter preferences for welfare which we combine with Current Population Survey data to determine the voter in each state who has the median preferred welfare benefit level. Our analysis reveals considerable evidence in support of a role for declining real wages in the decline of welfare benefits.
Author Keywords: Welfare benefits; USA; Wage inequality

Article:

1. Introduction

Welfare benefits in the United States paid to persons through the Aid to Families with Dependent Children (AFDC) program, the main cash assistance program for the poor from 1935 through 1997, declined in real terms from the late 1960s through the 1990s. State legislatures, which set nominal benefit levels, failed to raise those levels to a degree sufficient to keep up with inflation, and in recent years sometimes even reduced nominal benefit levels. Other major welfare programs – most notably the Food Stamp and Medicaid programs – grew in the early 1970s, causing the sum of all three benefits to rise, but even this benefit sum declined in real terms after the mid-1970s. While the rate of benefit decline slowed in the 1980s, and may even have flattened out, real benefits in the welfare system today are not much above what they were in 1970.

Considerable scholarly attention has been focused on the causes of this decline. The "demand" for welfare benefits has been treated with a public good model at least since Orr (1976), which, in its simplest variant, implies that income and price should determine demand. But real income in the aggregate and in all states has risen over the last 25 years, albeit at a slower rate than in earlier periods; and none of the variables affecting the tax price of state welfare benefits (federal matching rates, the AFDC caseload itself, etc.) have changed enough, or in a consistent fashion throughout the period, to provide a satisfactory explanation. The hypothesis that the federally-funded Food Stamp program may have "displaced" the AFDC program has also been studied (Orr, 1979 and Gramlich, 1982 and Moffitt, 1990a) with contradictory results across studies. Another hypothesis is that non-AFDC welfare expenditures, especially Medicaid, have risen rapidly and have crowded out AFDC expenditures. The results in the literature to date indicate, at best, weak evidence for this hypothesis as well (Moffitt, 1990a and Ribar and Wilhelm, 1994).

The hypothesis that we explore in this paper is related to the increase in wage inequality in the U.S., which has been the subject of much attention in recent years (see Levy and Murnane, 1992, for a review of the early studies). An oft-noted, and important feature of that trend, is a decline not only in wage rates in the lower tail of the distribution relative to higher-skilled groups, but also an absolute decline in lower-tail real wages. If voters wish to maintain a target benefit–wage ratio, either for reasons of equity with the nonwelfare working poor or to control work disincentives, a fall in the low-skill wage could lead voters to reduce welfare benefit levels. This stands in contrast to the standard view that a reduction in unskilled wages, by raising poverty rates, leads voters to increase benefit levels.

We test this hypothesis with a panel of U.S. state-level data over the period 1969–1992 with information on both benefits and low-skilled wages. Our preferred test uses now-conventional state fixed effects estimation, where estimates are based on the correlation between changes over time in welfare benefits, on the one hand, and changes over time in low-skilled wages, on the other, across different states. We find considerable evidence that the drift in wages has played a contributory role. At the state level, year-to-year changes in benefits and low-skill wages are positively correlated across states, both unconditionally and conditional on a set of other regressors including controls for income and price effects. In addition, long-term trends in benefits and wages (after smoothing out fluctuations) are also positively related across states.

An additional feature of our analysis relates to the median voter model which underlies our analysis and most of those in the literature. While we take the median voter model as a maintained hypothesis in our work, we use data on individual preferences for welfare to circumvent the aggregation problem that arises when only state-level data on median income and population characteristics are available, a problem discussed by Bergstrom and Goodman (1973). Using a question on welfare preferences from a nationwide survey (the General Social Survey), we are able to estimate the income and socioeconomic characteristics of the median-preference voter in each state and use those as regression determinants of the state benefit level, thereby freeing the analysis from the very restrictive assumptions given by Bergstrom and Goodman necessary to justify the use of, and interpret the results from using, aggregate data.

The next section of the paper briefly describes trends in welfare benefits and related variables. Following that, we discuss conceptual aspects of the hypothesis in more detail. After a discussion of modeling and data considerations, we present our empirical results. A summary concludes the paper.

2. Trends in welfare benefits in the U.S.

That welfare benefits have declined markedly over the last twenty years is well-known. Fig. 1 shows the trend in the sum of the real AFDC and Food Stamp guarantees, the maximum payments for a family of four with no other income, over the period 1970–1993. The essentially monotonic downward trend reflects a more or less constant real Food Stamp benefit (Food Stamps are indexed to inflation) and a nominal AFDC benefit which has been rising slower than the rate of inflation. In the early 1990s, some states also reduced their nominal AFDC benefits. Fig. 2 shows the time trend in the real weekly wage (1987 dollars) measured at the 25th percentile of the distribution.[1] The figure shows that this measure of the unskilled wage declined for most of the period, although with fluctuations around trend, until the late 1980s, after which it has risen slightly. Other figures (not shown) show similar declines for the wages of high school graduates and at the 10th percentile of the wage distribution.

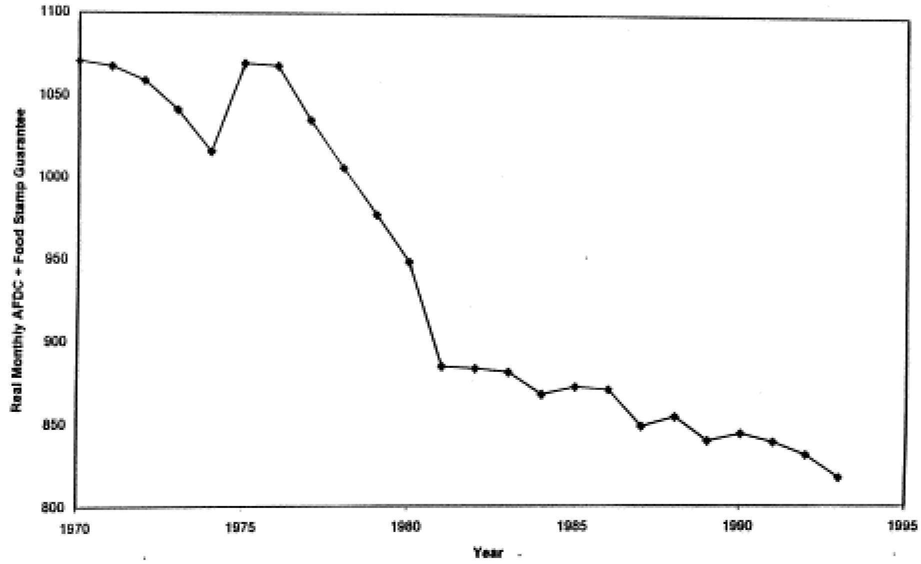

Fig. 1. Real monthly AFDC+ food stamp guarantee 1970–1993.

Fig. 2. 25th percentile of real weekly wage distribution 1970–1993.

Fig. 3 shows the trend in the ratio of AFDC and Food Stamp benefits to the 25th-percentile wage over the period. While the ratio exhibits some fluctuation, it has fallen fairly steadily over time. However, the rate of decline is considerably reduced from that of the benefit alone in Fig. 1. We take this observation as the starting point for our analysis.

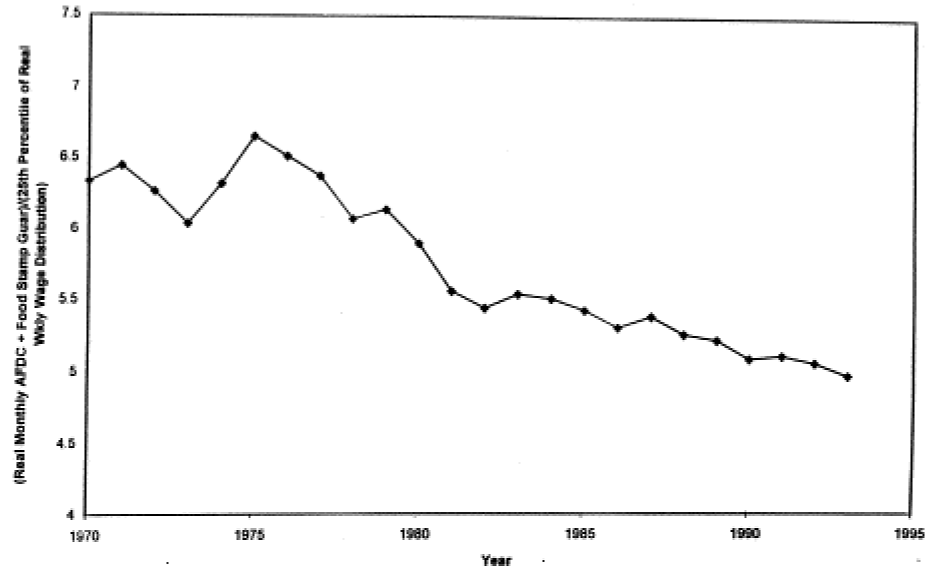

Fig. 3. Benefit wage ratio 1970–1993.

3. Modeling the influence of wages on benefits

There is an extensive public-economics literature on models of the state-level determination of AFDC benefits (Pauly, 1973; Orr, 1976; Orr, 1979; Gramlich, 1982; Plotnick and Winter, 1985; Moffitt, 1984; Moffitt, 1990a; Plotnick, 1986; Brown and Oates, 1987; Johnson, 1988; Shroder, 1995 and Ribar and Wilhelm, 1996, to name just a few). The standard model of redistribution from nonpoor to poor assumes that the existence of a group of poor persons imposes a negative externality on the utility of the nonpoor; optimal desires for redistribution then involve a balance between the marginal utility of redistribution and the income and efficiency losses attendant upon it (Atkinson, 1987). Because the AFDC program is largely a state-level program in the U.S., the literature on the determinants of AFDC benefits has translated the model of optimal redistribution to the local level, first theoretically by Pauly (1973) and then, in the first serious empirical application, by Orr (1976). The model assumes that each voter cares about his own income and, altruistically, about the income of a representative poor family, and that majority rule in a single-issue election leads to a benefit chosen by the median-preference voter. The other articles referenced above pursued theoretical and empirical variants on this theme. In the standard forms of these models, a decrease in the income of the poor leads to a increase in the marginal benefit of redistribution and hence to an increase in the equilibrium welfare benefit. We wish to posit models that could generate the opposite result.

We base our model, as is typical in the literature, on the utility function of the nonpoor voter on the assumption that poor voters are insufficiently large in number to generate a median voter. We consider three different types of variables affecting the utility of a representative nonpoor voter: consumption (C), hours of work (H), and other socioeconomic characteristics (X). Again following the standard formulation, we separate the arguments of the utility function into those of the individual himself and those of other individuals in the population. A general formulation of this utility function is:

$$(1)\quad U(C,\ \{C\}_{-1},\ H,\ \{H\}_{-1};\ X,\ \{X\}_{-1})$$

where C, H, and X denote quantities for the nonpoor individual himself and those variables in curly brackets with subscript −1 denote those variables for the residual set of individuals, presumed to include the target population for redistribution. It is important to note that a function as general as Eq. 1 can capture preferences for redistribution based not only on the absolute levels of C and H of the target population, but on their levels relative to those of the nonpoor individual himself. Thus if voters have a target level of inequality per se in mind, represented perhaps by a desired ratio of consumption, this can be captured in Eq. 1. “Relative income” theories of redistribution are of this type. Relatedly, the presence of the *X* characteristics is capable of representing not only the effect on redistributional preferences of the characteristics of the individual himself and those of the members of the target population, but also of the relationship between the two. This could matter if similarity or dissimilarity in characteristics between the donor and the recipient influence preferences, as in models of "social distance" (Kristov et al., 1992 and Cutler et al., 1993).

A stylized specialization of Eq. 1 can capture most of the important preference influences of interest here. Suppose that the target population can be considered to be composed of two homogeneous groups, one consisting of poor individuals receiving public assistance (A) and one consisting of poor individuals not receiving public assistance (NA). With additional separability assumptions between *C* and *H*, we have:

$$U = V\left(\underset{(+)}{C},\ \underset{(+)}{C_P^A}, \underset{(+)}{C_P^{NA}};\ X,\ X_P^A,\ X_P^{NA}\right) + Z\left(\underset{(+)}{H_P^A},\ \underset{(+)}{H_P^{NA}};\ H,\ X,\ X_P^A,\ X_P^{NA}\right) \quad (2)$$

where we have taken the labor supply of the nonpoor individual himself (*H*) to be exogenous. The budget constraint facing the individual is

$$(3)\ Y = Q(Y)B + C$$

$$(4)\ Q(Y) = k(Y)\,\frac{R}{N}$$

where *Y* is the income level of the individual, *B* is the level of the welfare benefit, *R* is the number of individuals on welfare (i.e., the caseload), *N* is the population in the state, and $k(Y)$ is a function nondecreasing in *Y* which represents the form of the tax function in the state. Thus $Q(Y)$ is the price of a one-dollar increase in *B* to an individual with income *Y*, and that price depends on the state per-capita caseload. We ignore federal subsidies for the time being.

To adequately capture the effects of benefits and unskilled wages on desired benefits, we add a set of functions describing the consumption, hours, and welfare-take-up decisions of poor individuals as follows, where *W* is the level of the wage for unskilled workers:

$$(5)\ C_P^A = c^A(\underset{(+)}{B}) \quad C_P^{NA} = c^{NA}(\underset{(+)}{W})$$

$$(6)\ H_P^A = h^A(\underset{(-)}{B}) \quad H_P^{NA} = h^{NA}(\underset{(+)}{W})$$

$$(7)\ R = R(\underset{(+)}{B}, \underset{(-)}{W}).$$

Given these functions, the optimal desired *B* can be obtained by taking the derivative of Eq. 2 w.r.t. *B* to obtain the marginal utility of increasing the benefit:

$$\text{(8)} \quad \text{MUB} = \frac{\partial U}{\partial B} = -V_1 Q(Y)(1+\epsilon) + V_2 \frac{\partial c^{\text{A}}}{\partial B} + Z_1 \frac{\partial h^{\text{A}}}{\partial B} = 0$$

where $\varepsilon=[(B/R)(\partial R/\partial B)]$ is the elasticity of the caseload w.r.t. the benefit. Optimal benefits are chosen by balancing the disutility of the reduction in consumption induced by the additional taxes needed to raise benefits, including any caseload increase that results, and the net utility of increasing the consumption of the poor at the expense of some reductions in their hours of work.

Although it is often presumed that higher income voters are more altruistic than middle-income voters, and that welfare benefits are in this sense a normal good, this does not imply that desired benefits are positively related to income. Taking the derivative of Eq. 8 w.r.t. Y, we have:

$$\text{(9)} \quad \frac{\partial \text{MUB}}{\partial Y} = -V_1 Q'(Y)(1+\epsilon) - [1 - Q'(Y)B]\left[V_{11} Q(Y)(1+\epsilon) - V_{21} \frac{\partial c^{\text{A}}}{\partial B}\right]$$

While it is true that the middle term in this expression implies (assuming that $[1-Q'(Y)B]$ is positive) that declining marginal utility of own consumption ($V_{11}<0$) leads to an increase in the desire for redistribution as income rises, the first term shows that an increase in tax payments as income rises may offset this. Less obvious is the third term, whose sign is determined by the sign of V_{21}, the effect of an increase in own consumption on the marginal utility of increasing the consumption of poor welfare recipients. Although there could be some type of “pure” positive income effect on redistribution, reflected in $V_{21}>0$, it is also possible that sympathy for the poor falls as “distance” from them (in income or consumption terms) grows; this would lead to $V_{21}<0$ and hence a decline in preferred benefits among those with higher income.

An implication of this point that is important for our empirical findings is that income effects may differ in sign depending on whether Y is increased only for the nonpoor voter in question, holding constant the income and consumption of the poor, or whether the entire income distribution shifts (e.g., through economic growth over time). If, for example, V_{21} is a negative function of the $C/C_{\text{P}}^{\text{A}}$ ratio, a shift upward in the entire distribution could generate a positive income effect whereas an increase in the income of one voter alone could generate a negative income effect. Empirical counterparts that would correspond to these different comparisons are, on the one hand, cross-sectional comparisons of benefit preferences of voters with different incomes – where absolute income effects cannot be separated from relative income effects – versus, on the other hand, comparisons over time of the benefit preferences of a population whose entire income distribution has shifted upward or downward.

The effect of the unskilled wage on benefit preferences is of key interest to us. Taking the partial derivative of Eq. 8 w.r.t. W, we have:

$$\frac{\partial \mathrm{MUB}}{\partial W} = -\left[V_1(1+\epsilon) - V_{11}BQ(Y)(1+\epsilon) + V_{21}B\frac{\partial c^{\mathrm{A}}}{\partial B}\right]\frac{\partial Q(Y)}{\partial W} - [V_1Q(Y)]\frac{\partial \epsilon}{\partial W} - \left[V_{13}Q(Y)(1+\epsilon) - V_{23}\frac{\partial c^{\mathrm{A}}}{\partial B}\right]\frac{\partial c^{\mathrm{NA}}}{\partial W} + \left[Z_{12}\frac{\partial h^{\mathrm{A}}}{\partial B}\right]\frac{\partial h^{\mathrm{NA}}}{\partial W} \quad (10)$$

The first two terms represent the effects of a change in W on the welfare caseload, which affects the tax price to the voter. These effects are ambiguous in sign because, although $\partial Q/\partial W<0$, the signs of $\partial\varepsilon/\partial W$ and V_{21} are not signable; the latter again represents relative income, social distance influences (in fact, a drop in poor consumption from a decline in W could increase social distance and lead to less desire for redistribution). The third and fourth terms represent effects working through the relative preferences of voters for the consumption and work hours of poor nonwelfare recipients. Of particular interest are the effects working through the cross utility terms V_{23} and Z_{12}, which represent the voter's relative preferences for welfare and nonwelfare consumption and work hours. If $V_{23}>0$, for example, as would arise if voters wish to maintain a target consumption ratio between the two poor groups, a decline in the wage will lead to a decline in the marginal utility of benefits. If $Z_{12}>0$, this could lead to the opposite result, however—if low wages cause the hours of work of nonwelfare recipients to decline, voters may be less bothered by the work disincentives experienced by welfare recipients.

Thus the model offers at least three reasons why benefits may fall if unskilled wages also fall: (i) falling wages induce greater caseloads and hence drive up the cost of a marginal increase in benefits; (ii) associated with the increase in caseload is a increase in work disincentives, which voters may dislike; and (iii) falling wages may create a gap between welfare and nonwelfare working poor which voters may wish to reduce by benefit reductions.

Turning to the preferences of the poor voters themselves, those preferences are, as noted earlier, unlikely to be decisive in voting; however, their preferences do matter because they help determine who the median voter is. Poor voters are likely to prefer higher welfare benefits for self-interest reasons, and may prefer them even if they are not on welfare because they might go on in the future; for them, benefits may play the role of insurance (Varian, 1980).[2] Consequently, we should expect preferences for benefits to fall with income at the lower ranges.

3.1. Aggregation of preferences

Now consider a model in which different voters $i=1,\ldots,N_s$ in different states $s=1,\ldots,S$ have preferences of the type we have just described. We can write a reduced-form benefit-demand function for individual i in state s as:

$$(11)\quad B_{is}* = g(Q_{is},\ Y_{is};\ X_{is},\ W_s)$$ where $B_{is}*$

is the desired benefit level, Q_{is} is the tax price, Y_{is} is income, X_{is} is a composite set of the X characteristics previously defined; and W_s is the level of the unskilled wage in the state. We expect price effects to be negative and income effects to be positive or negative, as described earlier, and we wish to test whether unskilled wages positively affect the benefit.

We take the median voter model, originally developed by Hotelling (1929), Bowen (1943), as a maintained hypothesis for our analysis. Although stable public choice under majority rule requires single-peaked preferences (which may not hold) as well as single-issue elections (whereas welfare benefits are chosen by state legislatures and governors who run in multiple-issue elections), it is unfortunately the case that equilibria in multiple issue elections are difficult to obtain except under very simple preference assumptions or without additional political structure imposed. We therefore leave progress on this front to future research. Thus, denoting the individual with median B_{is}* in a state by i=m, the actual welfare benefit in state s is:

$$(12)\ B_s = g(Q_{ms},\ Y_{ms};\ X_{ms},\ W_s).$$

Even with the maintained assumption of median voter choice, however, major estimation difficulties have arisen in the literature, some of which interact with our aim of estimating effects of state-level variables like the unskilled wage. The most serious difficulty is that neither Eq. 11 nor Eq. 12 can ordinarily be estimated because Eq. 11 requires information on individual preferences, which are generally unavailable, and because Eq. 12 requires knowledge of the characteristics of the median voter, which are also generally unknown because the median-preference voter, even in a micro data set, cannot be identified. Because the demand function in Eq. 11 contains multiple variables, the median of the dependent variable does not correspond in general to the median of any of the variables in the function, including income; the median of the dependent variable depends on the distribution of all of the variables and on the unknown coefficients which weight up those variables to a preference value.[3]

The conventional solution to this problem is to use aggregate, state-level variables in the estimation of Eq. 12 instead-specifically, the median income in the state, a tax price based on median income, and the mean population characteristics in place of X. Unfortunately, the conditions under which estimation of such an equation yields estimates of the function in Eq. 12 are extremely stringent, as demonstrated by Bergstrom and Goodman (1973). The most important restriction is that the income distributions within all subgroups defined by the X characteristics within a state must be proportional shifts of one another. The difficulty with this assumption in our case is that, if it fails, the unskilled wage in the state may proxy deviations in the income distribution from proportionality and hence deviations in the income of the median-preference person from median income. Hence any significant estimated effects of the unskilled wage may be spurious and may reflect influences of, rather than on, the median voter.[4]

Our approach to this problem is to first estimate conventional models using aggregate state-level data, which will be valid under the Bergstrom-Goodman conditions, but also to use a micro-level data set on welfare preferences to obtain the actual income and characteristics of the voter with median preferences. Welfare preference information is available from a survey, the General Social Survey (GSS), which asks direct questions about preferences for welfare spending. We use the GSS to estimate demand functions for individual preferences as a function of income and other characteristics, and then apply these estimated functions to the Current Population Survey (CPS) on a state-by-state basis to determine the median-preference individual and his or her characteristics (we use the CPS because the GSS sample sizes by state are too small for reliable estimation of the median). We then use these characteristics in Eq. 12 and compare the results to those of the aggregate model. This will furnish an implicit test of the Bergstrom-Goodman conditions as a whole, and will also provide us with an estimate of the effects of the unskilled wage not dependent on any of those conditions.[5]

4. Results: conventional aggregate model

For the aggregate analysis we employ a data set consisting of all U.S. states in each year from 1969 to 1992. Definitions and descriptive statistics for these variables are presented in Table 1. The dependent variable is the log of the real monthly AFDC benefit per recipient. We also test the AFDC guarantee (maximum amount payable to a family of four), which has been used in a fair number of previous studies. We prefer benefits per recipient because they more accurately reflect the influence of deductions and other benefit-formula factors not easily captured in the guarantee; on the other hand, the benefit captures some income and family-size variation as well. Fortunately, the two variables are highly positively correlated and our results, for the most part, do not depend on which measure is used.

Table 1. Variable definitions and descriptive statistics

Variable	Definition	Means	(Std. dev.)
State Level (*N*=1150)			
AFDC benefits per recipient	Total AFDC monthly benefits divided by recipients	119.22	(43.37)
AFDC guarantee for a family of four	Maximum monthly amount available to a family of four	495.13	(191.73)
Income per capita	Total personal income per capita	13227	(2556)
Price	Tax price of AFDC benefits under uniform taxation (Recipiency×State financing share)	0.018	(0.010)
Recipiency	Average monthly number of AFDC recipients divided by state population	0.040	(0.014)
State financing share	Share of AFDC expenditures paid by the state	0.448	(0.187)
Percent black	Percent of population of African descent	0.092	(0.091)
Percent under age 15	Percent of the population 14 years and under	0.244	(0.032)
Percent over age 64	Percent of the population 65 years and older	0.110	(0.023)
Percent with high school	Percent of the population with high school	0.662	(0.108)
Percent with college	Percent of the population with college or more	0.157	(0.044)
A.D.A. ranking	*American for Democratic Action* liberal rating of states' House and Senate delegations	0.430	(0.219)
State Level – from the CPS (N=1150)			
Low-skill wage	Weekly wage at the 25th percentile of the state wage distribution	156.66	(23.59)
Median income	Median household income	29322	(4360)
General Social Survey (N=16 339)			
Welfare preferences	Spending too much (wants less)	0.517	(0.500)
	Spending about the right amount	0.291	(0.454)
	Spending too little (wants more)	0.192	(0.394)
Income-amount	Annual family income	26412	(20186)
Income-refused	Refused to answer income question	0.034	(0.181)
Married	Respondent currently married	0.632	(0.482)
Education	High school	0.515	(0.500)
	Associate or junior college	0.032	(0.175)
	College	0.110	(0.312)
	Graduate	0.050	(0.217)
Black	Of African descent	0.104	(0.306)
Age	Years/100	0.445	(0.173)
Female	Female	0.540	(0.498)
Employed	Currently working	0.597	(0.490)
Unemployed	Currently unemployed	0.032	(0.175)
Hours worked	Weekly hours of work	24.249	(22.516)
Blue collar	Craftsmen, kindred workers, operatives (conditional on being employed)	0.180	(0.384)
Service worker	Clerical and kindred workers, service workers (conditional on being employed)	0.175	(0.380)
Farmer	Farmer or farm laborer (conditional on being employed)	0.013	(0.115)
Rural	Lives in an area of less than 2,500	0.183	(0.387)
Number of adults	Number of adults living in the household	2.005	(0.794)
Number of children	Number of children living in the household	0.921	(1.299)
Single mother	Respondent is single woman with children	0.063	(0.243)
Unemployed in last 10 years[a]	Had at least one unemployment spell during the last 10 years	0.292	(0.455)

Ever received government aid[a]	Ever received welfare, unemployment insurance, or other government aid	0.359	(0.480)
Women should be in home[a]	Thinks women should take care of their homes and leave running the country to men	0.286	(0.452)
Women can work[a]	Approves of a woman earning money even if her husband is capable of supporting her	0.747	(0.435)
Premarital sex is wrong[a]	Thinks premarital sex is wrong	0.393	(0.489)
Hard work is important[a]	Thinks hard work is most important for getting ahead	0.638	(0.481)
Luck is important[a]	Thinks luck or help from others is most important for getting ahead	0.120	(0.326)
Feel income is far below average[a]	Considers family income to be far below average	0.049	(0.215)
Feel income is below average[a]	Considers family income to be below average	0.229	(0.420)
Feel income is above average[a]	Considers family income to be above average	0.199	(0.399)
Democrat	Thinks of self as a Democrat	0.401	(0.490)
Independent	Thinks of self as an Independent	0.341	(0.474)
Republican	Thinks of self as a Republican	0.242	(0.428)
Liberal[a]	Thinks of self as liberal	0.279	(0.448)
Conservative[a]	Thinks of self as conservative	0.331	(0.471)

[a] The average omits additional missing data; additional data are missing because these questions were not included in each survey year.

The regressors in the equations include total personal income per capita in the state, the benefit price assuming uniform taxation and with federal matching (hence $k(Y)$=1 and the tax price is multiplied by one minus the federal matching rate), an unskilled wage variable in the state, and a set of demographic and political characteristics (percent black, percent aged, etc.).[6] The wage variable is computed from the March files of the CPS from 1969 to 1992. In each year, we compute the real average weekly wage (1987 dollars) of all workers in each state aged 16–64 and then select the wage at the 25th percentile of the distribution to proxy the wage for unskilled workers. We also compute other percentile points, the 10th percentile for example, and test them later in the paper. We use weekly wages rather than hourly wages because the CPS does not have a good measure of hours worked per week over the years prior to 1976. Past work has indicated, however, that hourly and weekly wages are highly correlated.[7]

Table 2 shows the results using the log of real AFDC benefits per recipient as the dependent variable. Column (1) includes only the income and price variables, as well as year dummies, and reveals significant positive and negative effects of each, respectively. Column (2) shows the effect of adding the unskilled wage, which has a significantly negative coefficient. However, when state-level fixed effects are added, as shown in column (3), the coefficient reverses in sign and becomes significantly positive.

The change of results with the introduction of state fixed effects is common in the literature and is unsurprising. It arises in the present case because many high-benefit states in the U.S. also have low real wages but, over time, states with more rapidly falling wages have experienced more rapidly falling benefits.[8] Since a specification test clearly rejects column (2), the implication is that the cross-sectional relationship is spuriously picking up fixed unobserved taste and other omitted state-level variables that are positively correlated with benefits but negatively correlated with wages.

The positive correlation between changes in benefits and wages over time is strong enough that it can be seen from a simple examination of the relationship between unconditional growth rates of benefits and wages by region, as shown in Table 3.[9] The table shows the change in average log benefits and log wages between two 11-year periods, 1969–1980 and 1981–1992 (our time period split in half). Log wages are measured relative to the median, to control for differences in the growth rate of the whole wage distribution and not just its lower portion. As the table indicates, benefits fell the most between the 1970s and 1980s in the mid-Atlantic region (New Jersey, New York, and Pennsylvania), and fell the least in the Pacific region (California, Alaska, Hawaii, and northwestern states). At the same time, inequality in the lower-tail of the wage distribution increased the most in the first group and increased less (in fact, decreased) for the latter group of states. Other groups fit this pattern as well: the industrial states of the Midwest (Michigan and Ohio, for example), experienced large increases in inequality and also reduced AFDC benefits by sizable amounts. The southeastern states in the U.S., which reduced benefits by very little compared to the rest of the country, also experienced little growth in inequality.[10]

Table 2. State-level regression analysis of determinants of AFDC benefits

Table 3. Trends in wage inequality and AFDC benefits by region, 1969–1980 to 1981–1992

Variable	(1)	(2)	(3)	(4)	(5)	(6)
Log	1.867 * **	2.202 * **	0.247 * **	0.090	0.218 * **	−0.016
(income per capita)	(0.064)	(0.077)	(0.070)	(0.087)	(0.073)	(0.100)
Log	−0.075 * **	−0.040 * *	−0.055 * **	−0.055 * **	−0.051 * **	−0.051 * *
(price)	(0.020)	(0.020)	(0.011)	(0.011)	(0.010)	(0.011)
Log	–	−0.594 * **	0.220 * **	0.228 * **	0.146 * **	–
(low-skill wage)		(0.078)	(0.039)	(0.039)	(0.035)	
Predicted log	–	–	–	–	–	0.365 * *
(low-skill wage)						(0.074)
Residual log	–	–	–	–	–	0.175 * *
(low-skill wage)						(0.046)
Percent black	–	–	–	−1.469 * **	−1.209	−1.465 * *
				(0.671)	(1.582)	(0.670)
Percent over age 64	–	–	–	−3.976 * **	−2.449 * *	−3.949 * *
				(0.863)	(1.180)	(0.861)
Percent under age 15	–	–	–	−3.625 * **	0.007	−3.481 * *
				(0.621)	(0.703)	(0.623)
Percent with high	–	–	–	−0.306	2.214 * **	−0.223
school				(0.266)	(0.727)	(0.269)
Percent with college	–	–	–	−0.652	−1.076	−0.550
				(0.482)	(1.454)	(0.483)
A.D.A. ranking	–	–	–	−0.036	−0.001	−0.039
				(0.027)	(0.024)	(0.027)
Time effects	Yes	Yes	Yes	Yes	Yes	Yes
State effects	No	No	Yes	Yes	Yes	Yes
State×trend effects	No	No	No	No	Yes	No
R^2	0.481	0.506	0.946	0.948	0.973	0.948

* Significant at 0.10 level.
** Significant at 0.05 level.
*** Significant at 0.01 level.
Note: Results based on 1969–1992 data from the 50 states (1150 observations). Dependent variable is the log of AFDC benefits per recipient. All of the independ
variables have been lagged one year. Standard errors appear in parentheses.

Region	Change in log real AFDC benefits per recipient	Change in log of the ratio of low-skill wage to median wage
Mid-Atlantic	−0.277	−0.032
East North Central	−0.194	−0.027
West South Central	−0.187	0.000
West North Central	−0.170	0.010
East South Central	−0.170	−0.010
Mountain	−0.132	0.009
South Atlantic	−0.038	−0.000
New England	−0.034	0.022
Pacific	−0.024	0.023

Note: The entries shown are changes in the 11-year averages between the two periods, taken over all states in each grouping.
Regions: Mid-Atlantic: New York, New Jersey, Pennsylvania; East North Central: Wisconsin, Illinois, Michigan, Indiana, Ohio; West South Central: Texas, Oklahoma, Arkansas, Louisiana; West North Central: North Dakota, South Dakota, Nebraska, Kansas, Minnesota, Iowa, Missouri; East South Central: Kentucky, Tennessee, Mississippi, Alabama; Mountain: Montana, Idaho, Wyoming, Colorado, Utah, Arizona, New Mexico, Nevada; South Atlantic: Florida, Georgia, South Carolina, North Carolina, Virginia, West Virginia, DC, Maryland, Delaware; New England: Vermont, New Hampshire, Maine, Massachusetts, Rhode Island, Connecticut; Pacific: California, Oregon, Washington, Alaska, Hawaii.

Returning to column (3) of Table 2, it can also be seen that, while price effects are not much affected by the introduction of state fixed effects, income effects fall markedly. Errors-in-variables in the income variable is one explanation for this reduction, but cross-sectional bias (high income states have high benefits) may also be at work. We will return to this issue when we use micro-level data on incomes. The fourth column in the table adds the demographic variables in the state. The coefficients on these variables are jointly significant. Their introduction further reduces the estimated income effect but has no effect on the wage coefficient.[11]

In columns (5) and (6) we test whether our results are robust to the presence of differential growth rates of benefits in different states by adding interaction variables between the state dummies and a time trend. With these state-trend fixed effects included, the coefficients on the other regressors in the equation reflect their relationships with year-to-year benefit movements apart from, or on top of, linear trends. The table shows that adding these effects changes a number of the coefficients but, while reducing the magnitude of the wage effect, leaves it positive and significant (the interactions are jointly significant). In column (6) the wage effect that works through the state-year trend is added back into the equation by decomposing the wage variable into two components, one equal to the value predicted from a first-stage regression of the wage on year, state, and state-trend dummies, and the other equal to the residual from this regression; the state-trend fixed effects are now omitted so that their effect can work through the predicted log wage variable. This procedure is equivalent to an instrumental variables procedure in which the state-trend dummies are used as instruments for the wage and generate consistent estimates under the alternative model in which state-specific year-to-year wage fluctuations around trend contain measurement error but there are no true state-trend fixed effects.[12] The results show a much larger positive effect of the predicted wage on benefits than for the actual wage – and, consistent with this, a smaller effect of the residual wage (the average effect in column (4) falls necessarily between the two estimates in column (6)). Thus we find that, indeed, states with larger trend growth rates of unskilled wages –

or, more accurately, slower declines – have also had slower trend declines in welfare benefit levels, confirming the results in Table 3 in a regression context.[13]

We conduct a number of additional sensitivity tests to this model and these results, including tests using different percentile points for both the low-skill wage and household income, using different measures of the wage, and testing for the effects of endogeneity in the tax price and other variables. However, we postpone a discussion of these tests until we use micro data to reformulate the median voter model.

5. Results: micro-based median voter model

In this section we report results using a micro-level survey on welfare preferences together with the CPS to determine the characteristics of the individual with median preferences in each state.

5.1. General social survey

The GSS is a nationally-representative, repeated cross-section survey which has been conducted every year since 1972 (except 1979, 1981, and 1992) and which interviews approximately 1500 individuals each year (Davis and Smith, 1992). Questions are asked about opinions and beliefs on a wide range of topics, and the question wording is intentionally kept fixed over time to be able to conduct meaningful time-series analyses. One of the questions concerns welfare spending. Respondents are asked whether they believe “too much,” “too little,” or the “right amount” is being spent on welfare. The exact question is shown in Table 4.[14]

Table 4. General social survey question on welfare preferences

Table 4

General social survey question on welfare preferences

We are faced with many problems in this country, none of which can be solved easily or inexpensively. I’m going to name some of these problems, and for each one I’d like you to tell me whether you think we’re spending too much money on it, too little money, or about the right amount. First (READ ITEM A) . . . are we spending too much, too little, or about the right amount on (ITEM)?

A.	Space exploration program
B.	Improving and protecting the environment
C.	Improving and protecting the nation’s health
D.	Solving the problems of the big cities
E.	Halting the rising crime rate
F.	Dealing with drug addiction
G.	Improving the nation’s education system
H.	Improving the condition of blacks
I.	The military, armaments, and defense
J.	Foreign aid
K.	Welfare
L.	Highways and Bridges
M.	Social Security
N.	Mass transportation
O.	Parks and recreation

Fig. 4 shows how the answers to these questions have changed over the period 1973–1993. There was a pronounced shift against welfare in the mid-1970s, just after the welfare caseload explosion of the late 1960s and early 1970s. Attitudes liberalized gradually over the latter half of the 1970s and early 1980s, leveled off in the mid-1980s, and have recently taken another turn against welfare. The means of these questions over the period are shown in Table 1, which indicates that about 52% of the population believed spending was too high, 19% believed spending was too low, and the residual 29% believed spending was at the right level.

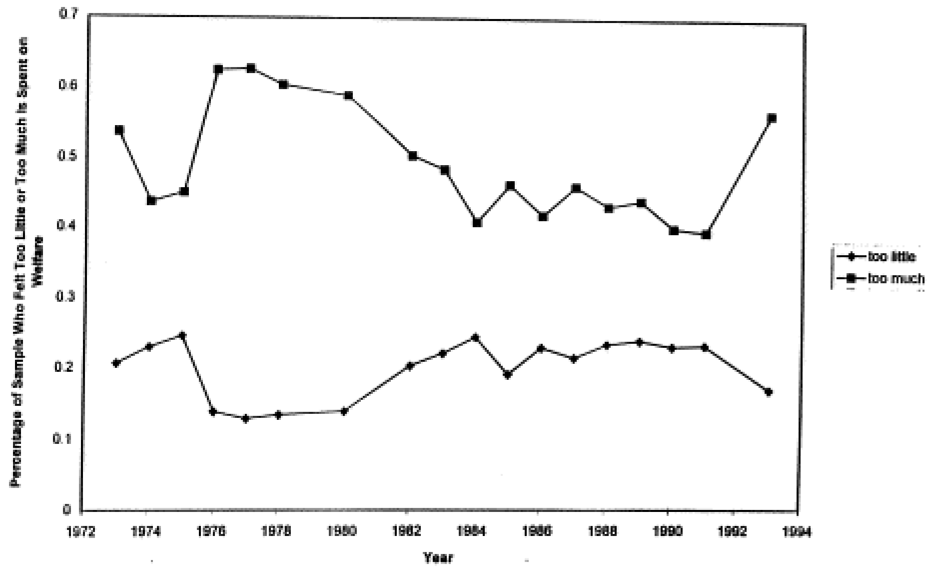

Fig. 4. Public opinion on welfare expenditure 1973–1993.

It should be noted that the percent of the sample wishing to lower spending exceeds 50% in some years according to Fig. 4 and nearly 52% of the sample in all years wished to do so, seemingly violating the median voter equilibrium assumption. However, these percentages are from the full national sample and hence are implicitly weighted by state population. When data by individual state in each year are examined, the proportions answering "too much" or "too little" rarely exceed 50%; the average fraction (across all states and years) answering "too much" is 47.3% and that answering "too little" is 21.5%. This suggests that the answers to the questions are being given relative to within-state factors; that is, that they are within-state preferences.[15]

Taking the answers to the GSS questions as proxies for B_{is}*, our primary goal is to determine the characteristics of the median-preference individual. The most straightforward approach would be to select the individual from the GSS in each state in each year who has the median value of the GSS question, but this is not possible both because the preference variable is categorical – and hence the median person cannot be uniquely identified from the individuals in the middle category replying that the "right amount" is being spent on welfare – and because the GSS sample sizes are too small to permit a reliable estimate by state and year. Our approach is, instead, to estimate a benefit-preference equation on the GSS as a function of income and other characteristics (i.e., the variables in Eq. 11) using the technique of ordered probit, which yields estimates of a continuous, latent preference index; and then to apply that estimated equation to the Current Population Survey (CPS) to select an individual in each state in each year who has the median value of the predicted index. The CPS has much larger sample sizes by state and year (generally no fewer than 300 or 400, and up to 6000) and hence will yield more reliable estimates.[16]

Estimates of the individual welfare preference models using the GSS data are shown in Table 5 (we pool the data over all years 1973–1990, giving us 16 339 observations). For the dependent variable we classify the answer "too much" spending on welfare as the lowest category on the scale and "too little" as the highest category so that the coefficients will represent effects on increased spending. The independent variables we include are listed in Table 1. Each specification contains a full set of state dummies interacted with year dummies. The inclusion of this full set of dummies allows us to obtain the effects of individual-level characteristics on preferences using entirely within-state variation, which, as we noted earlier, is the way the GSS question is answered.[17]

Table 5. Individual-level analysis of the determinants of welfare demand in the GSS

Variable	
Income ≤ \$20 000	−3.220 $***$
	(0.221)
Income ≥ \$20 000	−0.415 $***$
	(0.074)
Married	−0.120 $***$
	(0.025)
High school	−0.148 $***$
	(0.025)
Associate	−0.183 $***$
	(0.060)
College	0.016
	(0.039)
Graduate	0.298 $***$
	(0.052)
Black	0.707 $***$
	(0.033)
Age	−0.452
	(0.359)
Age-squared	−0.222
	(0.380)
Female	−0.056 $**$
	(0.023)
Employed	−0.100 $**$
	(0.050)
Unemployed	0.069
	(0.058)
Hours worked	−0.004 $***$
	(0.001)
Blue collar	0.057

	(0.035)
Service worker	0.048
	(0.034)
Farmer	−0.165*
	(0.097)
Rural	−0.095 * **
	(0.031)
Number of adults	0.097 * **
	(0.013)
Number of children	0.027 * **
	(0.009)
Single mother	0.071
	(0.045)
Income-refused	−0.696 * **
	(0.064)
Upper threshold	0.943 * **
	(0.012)
Log-likelihood	−15 094.2

* Significant at 0.10 level.
** Significant at 0.05 level.
*** Significant at 0.01 level.
Note: Maximum likelihood estimates of ordered probit models of welfare demand. The data are from the GSS 1973–1990. The functional form of income is a linear spline with a single knot placed at $20 000; the models are estimated with income expressed in $100 000s. The upper threshold is an estimate of the level of the latent index above which respondents answer that they want more welfare; the threshold for the latent index below which less welfare is demanded is normalized to zero. All equations include a full set of year*state dummies. The log-likelihood of the model with slopes restricted to zero is −16 618. Asymptotic standard errors appear in parentheses. N= 16 339.

In our first specification we restrict ourselves to variables that are in the GSS but also in the CPS. Price is omitted because, within-state, it has no variation if uniform taxation is assumed and is a linear function of income if proportional income taxation is assumed, in which case the income coefficient will absorb its effects.

The first two rows in the table show a spline for family income, which is shown to have a negative effect on benefit preferences, an effect which tapers off with income. The spline break point of $20 000 is chosen on the basis of Fig. 5, which plots the GSS questions versus income. Apart from normal sampling variance, which induces some fluctuations in the relationship, the figure shows that support for welfare benefits declines rapidly initially and then levels off as income rises. The negative relationship between income and benefit preferences clearly supports a self-interest interpretation of preferences. The breakeven level of the AFDC income (i.e., the eligibility income level) is not far from $20 000 in many states, which further supports this interpretation. The negative income result is also consistent with a "social distance" interpretation of preferences for redistribution, whereby those farther away, on the income scale, from the poor are less able to empathize with the poor. Within the context of the social distance interpretation, the failure of support for benefits to turn up at higher incomes signals that altruistic motives are insufficiently strong to offset the combination of increased social distance and higher tax prices.

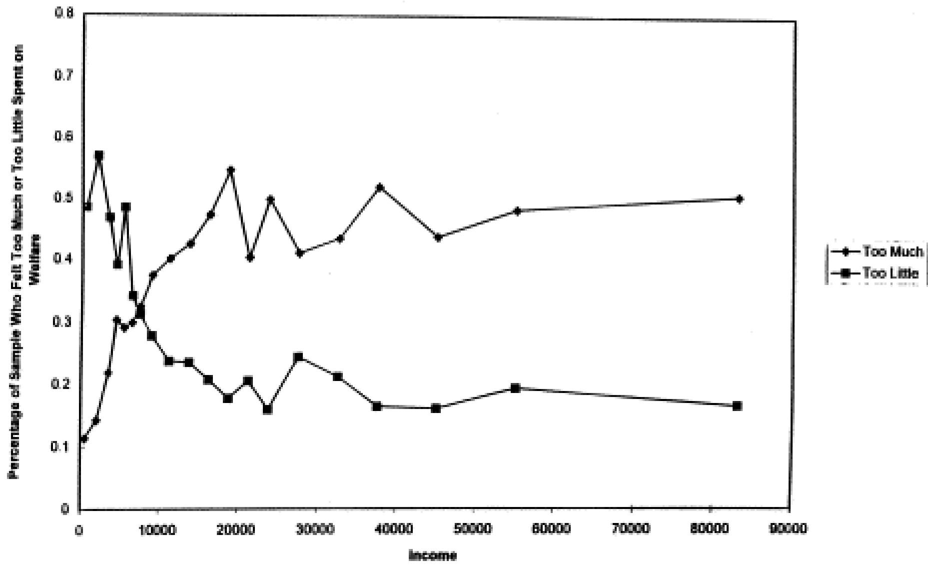

Fig. 5. Public opinion on welfare spending by income 1986–1990.

The negative income coefficients found here are not inconsistent with the positive income effects found in the aggregate analysis in part because the GSS coefficients incorporate price effects, but also because they do not measure the same thing as those in the aggregate analysis. The income coefficients in the GSS reflect the fact that higher income individuals have different preferences for redistribution, perhaps for the relative income or social distance reasons we have suggested; but this is not the same as increasing the income of a fixed individual, holding that individual's preferences fixed. The latter can be positive while the former is negative. This is, in econometric terms, a case of a conventional individual fixed effects model where the omitted fixed effect in a single cross-section, which may represent relative income, is negatively correlated with income but cancels out in a fixed effects, or first-difference, model.[18, 19]

The other coefficients in column (1) are also of interest because they imply a mixture of altruistic, social distance, and self-interest motivations for benefit preference. That higher benefits are preferred by those with larger family sizes, who are unmarried, who are black, and who do less work all suggest self-interest motives because these variables are all positively correlated with the probability of being on AFDC (a social distance interpretation also could be given to these results).[20] On the other hand, that higher benefits are preferred by those with higher education and by men is counter to self-interest explanations, and at least the education effect might reflect altruistic preferences acquired through schooling.

It is of interest to compare the effects of the individual sociodemographic characteristics here with those in Table 2, even though the comparison is complicated by the same considerations we discussed for income (namely, the coefficients here may be picking up preferences rather than their effects holding preferences fixed). For example, percent black is strongly negative in Table 2 but positive in Table 5. One explanation for this difference is that the Table 2 variable is a proxy for a higher black representation in the welfare caseload, and a consequent negative attitude by white voters, while the Table 5 variable is proxying attitudes toward welfare by black voters.[21] The college and age variables also have different signs in the two Tables. This should generate some concern that the state-level demographic variables in Table 2 are spuriously measuring factors other than voting preferences.

Table 6 shows the coefficients on several additional variables added to the model. These variables are in the GSS but not in the CPS; nevertheless, they are informative and of independent interest because they consider a broader range of potential influences on welfare preferences.[22] The estimates in column 1 show that those who were unemployed or who received government assistance in the past are more supportive of welfare, consistent with self-interest or social distance motives.[23] Column 2 adds variables which indicate respondents' family values, opinions about the rewards from work, and views of own income relative to the average. Those with more conservative opinions regarding the role of women in the home and premarital sex are less supportive of welfare. Respondents who think that “lucky breaks or help from others”, as opposed to hard work, are most important in getting ahead, are more supportive of welfare. Those who consider their income to be “far below average”, and to a lesser extent those who consider their income to be “below average”, are more supportive of welfare than those with average or better-than-average incomes. This last result is consistent with our argument that redistribution preferences depend upon relative position in the income distribution. The estimated effects of the political variables in column 3 are not surprising. Democratic party affiliation and politically liberal views are positively correlated with redistribution preferences. Independents support welfare more strongly than Republicans (the omitted category is those claiming another, or no, party affiliation) but not as strongly as Democrats.

Table 6. Individual-level analysis of the determinants of welfare demand in the GSS: additional variables

Variable	(1)	(2)	(3)
Unemployed in last 10 years	0.095 * **	0.079 * **	0.068 * **
	(0.026)	(0.026)	(0.026)
Ever received government aid	0.137 * **	0.128 * **	0.122 * **
	(0.026)	(0.026)	(0.026)
Women should be in home	–	−0.061 * *	−0.046
		(0.030)	(0.030)
Women can work	–	0.019	0.015
		(0.030)	(0.030)
Premarital sex is wrong	–	−0.125 * **	−0.085 * **
		(0.028)	(0.027)
Hard work is important	–	−0.012	−0.007
		(0.029)	(0.030)
Luck is important	–	0.086 * *	0.072*
		(0.042)	(0.043)
Feel income is far below average	–	0.248 * **	0.239 * **
		(0.047)	(0.047)
Feel income is below average	-	0.117 * **	0.118 * **
		(0.026)	(0.026)
Feel income is above average	–	−0.027	−0.020
		(0.030)	(0.030)
Democrat	–	–	0.318 * **
			(0.079)
Independent	–	–	0.200 * *
			(0.079)
Republican	–	–	0.089
			(0.080)
Liberal	–	–	0.206 * **
			(0.027)
Conservative	–	–	−0.122 * **
			(0.026)
Log-likelihood	−15 067.6	−15 022.1	−14 887.4

Note: See notes to . Equations include all other variables shown in .

.2. State-level regressions using median characteristics

Table 7 shows estimates of state-level models identical in all respects to that in Table 2, column (4) except that per capita income is replaced by various forms of median income or the income of the individual with median welfare preferences. The final column also includes the other characteristics of the median-preference individual. To determine the income and other characteristics of that individual, we apply the estimates from Table 5 to the March files of the CPS from 1969 to 1992 and order the individuals in the CPS in each state in each year by their predicted benefit-preference index from the ordered probit. We then select the individual in the median of this distribution.[24] Initial inspection of the data revealed that there was considerable fluctuation of individual characteristics in the neighborhood of the median-preference individual; that this could cause difficulties was confirmed by initial testing of these variables in the state-level regressions, which showed all coefficients to be insignificant. Therefore, we instead selected the central 10% of the preference distribution (i.e., all individuals 5% above and below the median) and averaged their characteristics for use in the state-level regressions. As in column (4) of Table 2, each specification includes year and state fixed effects and state demographics.

Table 7. State-level regression analysis of the determinants of AFDC benefits: alternative measures of median voter's income

Variable	Median income	Income of Individuals with median preferences	Income of individuals with median preferences adjusted for voting	Income and other characteristics of individuals with median preferences
Log (income)	0.106* (0.063)	0.039 (0.048)	0.042 (0.048)	0.055 (0.062)
Log (price)	−0.057*** (0.011)	−0.057*** (0.011)	−0.057*** (0.011)	−0.054*** (0.011)
Log (low-skill wage)	0.195*** (0.046)	0.227*** (0.041)	0.226*** (0.041)	0.264*** (0.048)
R^2	0.948	0.948	0.948	0.949

* Significant at 0.10 level.
** Significant at 0.05 level.
*** Significant at 0.01 level.
Note: Results based on 1969–1992 data from the 50 states (1150 observations). Dependent variable is the log of AFDC benefits per recipient. Regressions include state demographics and controls for time- and state-specific effects. The specification in column 4 also includes other characteristics of the median-preference person. All of the independent variables have been lagged one year. Standard errors appear in parentheses.

Column (1) of Table 7 simply replaces state-level mean income by the state median income taken from the CPS, an accurate application of the median voter theorem if the median-preference individual is in fact the person with median income. The income coefficient is approximately the same as in Table 2 but is more significant, while the price and wage coefficients are somewhat more different in magnitude but retain the same level of significance. The major conclusion we draw from this result is that the use of mean rather than median income does not make much of a difference and, most important for our purposes, does not affect the estimated influence of the unskilled wage on benefits.

In the second column we replace the median income variable by the income of the median-preference individual. The point estimate of the income effect is lower than in column (1) and loses significance, although the two are not significantly different because of the fairly large standard errors. The low-skill wage and price coefficients are now even closer to their counterparts in column 4 of Table 2. The interesting similarity between these results and those in column (1) arises in part because we found benefit preferences to be monotonic in income and in part because the other individual characteristics that have influenced the determination of the median-preference individual are fairly highly correlated with income.

In column (3) we make an attempt to adjust for the fact that not all individuals vote. Since voting probabilities are correlated with income and other characteristics, the individual with median preferences is not exactly the same as the voter with median preferences. We use a question in the GSS asking respondents whether they voted in the most recent Presidential election to estimate probit equations for the probability of voting, as a function of the same variables in our welfare preference equation; we then apply that equation to the individuals in the CPS to obtain predicted voting probabilities, which we then use with the individual-specific predicted welfare preferences to calculate a weighted median-preference individual. Column (3) shows the results using the average income of those within the 10% band of that person. The similarity of these results to those in column (2) shows our results to be robust to the incorporation of voter selectivity.

In the final column we return to the specification with the income of the median-preference individual but now also add the associated individual-level socioeconomic characteristics of that person (coefficients not shown). The income coefficient remains smaller in magnitude than that in Table 2 but is also insignificant, and the price and wage coefficients are not much affected by this expanded specification.

To conclude, therefore, we find that our exercise in applying a more accurate version of the median voter model using individual-level preference data does not affect the results on the main issue in our paper, namely, the estimated effect of the unskilled wage on observed benefits. The distortions introduced by using state-level median income and state-level mean demographics do not significantly affect the relationship of the wage to the benefit level. However, the analysis of the GSS has provided some interesting evidence on the individual-level correlates of welfare preferences in general which may be useful in future research on other aspects of this topic.

6. Sensitivity tests

Finally, we report the results of conducting a variety of sensitivity tests to the model, many of which we have referred to in passing. The sensitivity tests often affect some estimated parameters of our model but, as we will conclude, they show the coefficient on the unskilled wage, which is our major concern, to be quite robust.
We conduct our tests on the model specified in the last column of Table 7 but we show only the low-skill wage coefficient for brevity. The first test we perform uses the AFDC guarantee for a family of four as the dependent variable rather than AFDC benefits per person. As Table 8 indicates, this reduces the wage effect slightly, but it remains significant and of the same general magnitude as in Table 7.

Table 8. State-level regression analysis of the determinants of AFDC benefits: sensitivity analysis of the effect of low-skill wages

Change from baseline	Low-skill wage coefficient
Use log (AFDC guarantee) as the dependent variable, replacing log (benefits per recipient).	0.191 * ** (0.051)
Add household income at the 75th percentile as an independent variable.	0.231 * ** (0.050)
Use weekly wages at the 10th percentile as the measure of low-skill wages, replacing weekly wages at the 25th percentile.	0.135 * ** (0.029)
Use hourly wages at the 25th percentile as the measure of low-skill wages, replacing weekly wages at the 25th percentile.	0.268 * ** (0.071)
Use weekly wages of those with less than a high school education as the measure of low-skill wages, replacing weekly wages at the 25th percentile.	0.098 * ** (0.033)
Use household income at the 25th percentile as the measure of low-skill wages, replacing weekly wages at the 25th percentile.	0.161 * ** (0.057)
Add a state-level cost-of-living index as an independent variable.[a]	0.110* (0.058)
Use women's weekly wages at the 25th percentile as the measure of low-skill wages, replacing (men's and women's) weekly wages at the 25th percentile.	0.151 * ** (0.034)
Use instrumental variables (the ratio of men to women aged 15–44 and the state financing share) for price.	0.249 * ** (0.046)
Remove recipiency from the definition of price, leaving only the state financing share (this is a reduced-form).	0.298 * ** (0.047)
Model the error term as AR(1).	0.189 * ** (0.037)

[a]Regression includes the years 1981–90 only.
Notes: See the notes to .

Next, we add the 75th percentile of household income in the state to the equation both to crudely test the median voter assumption (because other percentile points should be insignificant) and also to test whether the unskilled wage, because it may reflect a general increase in variance, is partly picking up the effects of increased dispersion in the upper tail of the income distribution. This change also has little effect on the estimated low-skill wage coefficient. However, the coefficient on the 75th-percentile income variable is itself quite significant, inconsistent with the median voter assumption (coefficient=0.168, s.e.=0.084).

We next conduct several tests using alternative measures of the low-skill wage. We test the 10th percentile wage rather than the 25th percentile; the real hourly wage instead of the real weekly wage; the wage for high-school dropouts; and even household income at the 25th percentile rather than the individual wage. We also test whether using a state-level cost-of-living index (McMahon, 1991) to adjust for differences in cost of living affects the results.[25] As Table 8 shows, these variations generate a range of wage estimates of 0.098–0.268, which are generally somewhat lower than the coefficient in Table 7. However, all effects remain positive and highly significant.

We next test the effect of using the female weekly wage instead of the overall wage as a simple test of the potential for endogeneity of the wage in the model. To the extent that welfare benefit levels affect labor supply of women, the equilibrium wage in the state labor market may be partly a result of benefits rather than their cause. In particular, higher benefits may reduce labor supply and increase the wage, leading to a spuriously positive relationship. If this effect is present, we should expect the wage coefficient to be larger when the female wage rather than the overall wage is used because over 90% of welfare adults are women. As Table 8 shows, the opposite appears to be the case – the coefficient actually falls when the female wage is used (though it remains positive and significant). We take this is as an indication that significant feedback effects on the wage are unlikely to be present.[26]

We next conduct two tests for possible endogeneity of the state tax price. Since that price is a function of the caseload (although lagged), which is no doubt responsive to benefits, the potential for endogeneity is clear. This issue has been heavily discussed in the literature on AFDC benefit determination at the state level (see Ribar and Wilhelm, 1996, for a recent review). We first provide instrumental variable estimates using the ratio of men to women aged 15–44 (the sex ratio should affect recipiency rates) and the state financing share as instruments for price.[27] In addition, to avoid having to make assumptions about legitimate instruments, we simply estimate a reduced form version by removing the recipiency rate from the price definition altogether, leaving only the state financing share.[28] If recipiency is endogenous and if the consequent bias affects the wage coefficient as well as the price coefficient, the wage coefficient should be significantly altered in this specification. As Table 8 indicates, however, neither of these tests has much of an effect on the wage coefficient at all. Therefore we find no evidence of price endogeneity problems that are connected to our estimated wage effects.

Finally, we relax the restrictions in the simple fixed effects version of our cross-state, over-time panel model by introducing serial correlation in the error term through an AR(1) specification, thus accounting for within-state correlation. This introduction reduces the magnitude of the low-skill wage effect slightly, but it remains positive and highly significant.

7. Conclusions

In this paper we have tested whether the decline in welfare benefits in the U.S. over the last twenty five years can be partly attributed to reductions in the wage rates of unskilled workers at the bottom of the wage distribution. Our results, based on conventional state fixed effects models, provide considerable support for the hypothesis. The result is strong enough that it can be seen from unconditional relationships between changes in unskilled wages and AFDC benefits across states from the 1970s to the 1980s, and it is robust to the inclusion of a number of other variables.

In addition, our analysis illustrated a new method of implementing median voter models when direct individual-level information on preferences is available. We used those data to determine who the median preference individual in the jurisdiction actually is to avoid some of the difficulties attendant upon using aggregate-level income and other characteristics to proxy those of the median individual. Our wage effects on benefits remain positive in this exercise and in additional sensitivity testing of the model as well.

There are several areas in which further work would be fruitful. For instance, the availability of individual preference data opens the possibility of estimating alternatives to the median voter model, whereas in our paper, and in the previous literature, the median voter model is a maintained hypothesis. However, the available preference data have limitations themselves. While our results suggest that self-interest or altruism filtered through social distance influences support for welfare, additional data are necessary to distinguish between these motives. Such data could also be gathered for the purpose of discerning which of the voter concerns (work disincentives or equity) we have discussed is the most important factor underlying the low-skill wage effect. Alternatively, this question might be approached from a reexamination of the low-skill wage effect in the years following the 1996 welfare reform which should serve to mitigate voter concern about welfare's work disincentives.

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Appendix

Variable	Source
AFDC benefits per recipient	*Social Security Bulletin*
AFDC guarantee-family of four	Unpublished data from the Office of Family Assistance
Income per capita	*Regional Economic Information System (REIS)*
Recipiency	Recipiency data from *Social Security Bulletin*, Quarterly and Annual *Public Assistance Statistics*, and *Green Book*; populatio data from *REIS*
State financing share	Data from *Characteristics of State Plans for AFDC*, *Social Security Bulletin*, *Green Book*, and unpublished tables from Administration for Children and Families
Percent black	*Decennial Census, Intercensal Estimates*
Percent over age 64	*Decennial Census, Intercensal Estimates*
Percent under age 15	*Decennial Census, Intercensal Estimates*
Percent with high school	*Decennial Census* (intercensal figures interpolated)
Percent with college	*Decennial Census* (intercensal figures interpolated)
A.D.A. ranking	*A.D.A. Today* (newsletter of *Americans for Democratic Action*)
Ratio of men-to-women, aged 15–44	*Decennial Census, Intercensal Estimates*

Notes

[1]We employ the March Current Population Survey (CPS) in each year to obtain these figures. We use all workers, regardless of age, sex, race, or household headship status, and we divide annual earnings by annual weeks of work. The series is deflated with the personal consumption expenditure deflator from the National Income and Product Accounts.

[2]An insurance motive may extend up into the income distribution as well, although presumably not farther than the variance of individual transitory income warrants and not for groups (like married couples) who are unlikely to be eligible for welfare.

[3]The median of a sum is asymptotically equal to the sum of the medians of the components if the components are independently and symmetrically distributed (Bergstrom and Goodman, 1973, p. 294). But income is correlated with the X characteristics, violating the independence assumption, at a minimum. Note that the wage variable in Eq. 11 is statewide and hence has no effect on the within-state preference distribution, and that the tax-price variable has within-state variation only because of variation in Y and hence can be folded into income.

Thus it is the presence of the *X* characteristics that represents the most important difficulty in mapping preferences into income.

[4]It is important to note that the Bergstrom-Goodman result does not imply that the median-income voter is the median-preference voter; instead, the authors show that the demand of the median-preference voter is proportional to the median income in the state (under their assumptions), which is not necessarily the income of that voter. It should also be noted that the coefficients on the aggregate *X* variables in a state-level equation are not the same as the coefficients on the individual *X* variables in Eq. 12; the aggregate *X* variables go into the proportionality factor shifting median income to the median preference person. See Bergstrom and Goodman, 1973 (pp. 286–287, 295–296).

[5]A few other papers in the literature (e.g., Bergstrom et al., 1982 and Husted, 1989) have used micro-level preference data as well. However, these authors stopped after estimating individual micro-level demand equations; they made no attempt to use their estimates to determine the median voter or to test whether the characteristics of the implied median voter could successfully explain the choice of the local public good.

[6]We begin with mean state income, rather than median, to follow the previous literature which has used mean state income instead of the income of the median-preference person. This is another source of disjuncture with the theory which we can treat more accurately with the micro data. We should also note that we tested in the tax price variable an assumption of proportional taxation, in which case the variable is multiplied by the ratio of individual income to state mean income, with little change in the results.

[7]In our data, the 25th percentile hourly wage and the 25th percentile weekly wage have a correlation coefficient of 0.61. The correlation coefficient between the hourly and weekly wages of high school graduates is 0.96.

[8]The fixed effects are jointly highly significant and the two wage coefficients are clearly significantly different from each other, so there is little question that a formal specification test would reject the no-fixed-effects model.

[9]It is easier to demonstrate the effect with regions than with states but it holds for the latter as well.

[10]This regional pattern of inequality trends is very close to that found by Topel (1994). He also found that the industrial states of the Midwest and Mid-Atlantic have experienced the largest increases in wage inequality.

[11]We might note at this point that the time dummies in all of our models remain highly significant, although reduced in average magnitude by the inclusion of the wage. Thus the wage is a contributing factor to the benefit decline but does not explain it all.

[12]At issue is whether it is better to purge the wage effect of possibly biasing state-trend fixed effects, as in column (5), or to follow the diametrically opposite procedure of purging the wage effect of its year-to-year fluctuations around trend because they may contain measurement error. We do not provide tests to discriminate between the two but simply present the results and leave the choice to the reader; both generate positive wage coefficients.

[13]By the principle of auxiliary regression estimation of linear regression coefficients, the coefficient on the residual in column (6), 0.175, should be identical to the wage coefficient in column (5). It is not because we have not included all of the other regressors in our first stage equation.

[14]The welfare question is known for wording experiments in which responses shift dramatically if “welfare” is replaced by “assistance to” or “caring for” the poor (Heclo, 1986); we do not use either of these latter two questions. A question order experiment in 1976 placed a tax question before the battery of spending questions for half the respondents. This had no effect on average responses to the welfare question in the two groups of respondents, and therefore we use both groups in our analysis.

[15]The GSS question might be interpreted as asking for preferences relative to national welfare spending because so many of the other items in the list are national in nature (see Table 4). However, if most respondents were answering in that fashion, the majority of respondents in high-benefit states should be expected to wish to increase spending and the majority of respondents in low-benefit states should wish to decrease spending. In fact, not only is this not the case in the data, the fractions wishing to decrease benefits are slightly smaller, not greater, in the low-benefit states compared to the high-benefit states.

[16]We do not claim originality for the idea of estimating the coefficients on a continuous latent index from categorical spending preference equations; Bergstrom et al. (1982) did this as well, using ordered logit.

[17]This is a conservative approach because, if state and national variables affect the answers to the GSS questions either by affecting the reference point or by directly affecting preferences conditional on that point, those effects

will be absorbed by our dummy variables and will leave the coefficients on the individual variables valid and unbiased estimates of within-state rankings, which is all that is needed to identify the median preference individual (which is the primary object of this exercise).

[18]An individual fixed-effects model is not the same as a state fixed-effects model but, as pointed out in prior work (Moffitt, 1993), the two models are equivalent asymptotically if births, deaths, and migration are ignored. If the population in a state is unchanging over time, variable means computed from a series of independent, repeated samples of the state population are consistent estimates of the means that would be provided by true panel data because the individuals sampled from are always the same. Consequently, a state fixed effects model would yield consistent estimates of an individual fixed effects model.

[19]Table 2 shows positive income effects even without state fixed effects, but this is still an across-state relationship and not the type of within-state relationship that the GSS provides.

[20]Unfortunately, the GSS usually did not ask welfare participation questions and hence we cannot, for example, estimate a "probability of participation" index in the GSS and test that as a regressor.

[21]However, it should be noted, as we did previously, that the Bergstrom-Goodman result does not imply that the coefficients on aggregate *X* variables reflect the effects of the median voter's *X*. In addition, the Bergstrom-Goodman interpretation of the role of aggregate *X* variables in a state-level equation breaks down completely if those *X* variables measure the characteristics of the recipient rather than the donor. See also Poterba (forthcoming) for a study of the effect of demographics on public choice.

[22]These variables are regularly collected in the GSS but, with the exception of political party identification, have not appeared in every survey year. Therefore, the models include a set of (unreported) dummies for each of these variables indicating observations with missing data.

[23]We attempted to discriminate between these two explanations by interacting current income with the indicator of past unemployment, on the presumption that high income individuals are unlikely to be motivated by self-interest because their probabilities of welfare participation are so low. The estimated effect of the interaction term is positive, favoring the social distance explanation, but falls short of statistical significance.

[24]We ignore the influence of the residual when making the prediction of the preference index. We appeal to Lemma 3 of Bergstrom and Goodman, 1973 (p. 294) that the median of a sum of independent symmetric distributions is the sum of the medians, which we assume will hold approximately; since the median of the residual is zero, the median preference will be the median of the predicted index.

[25]The low-skill wage effect could be an artifact of (i) co-movement between the low-skill wage and local cost-of-living and (ii) states partially adjusting benefits according to the latter.

[26]This is perhaps unsurprising because low-income female heads, even when off welfare, constitute only a tiny fraction of the state labor force. Also, Moffitt (1990b) conducted an extensive examination of whether the welfare system could have contributed to the observed growth in wage inequality in the U.S. (by depressing wages at the low end, from labor supply effects) and concluded that there was no evidence for such a contribution.

[27]Ribar and Wilhelm (1996) found that the ratio of men-to-women and the state financing share significantly predict the log price in the anticipated direction and that the state share survives a test for its over-identifying restriction.

[28]Recall that the price is entered in log form, so it is arithmetically equal to the sum of the log of the recipiency rate and the log of the state financing share. Taking the log of the recipiency rate as a function of the other regressors in the equation gives us an implied reduced form.